A New Empirical Method for Valuing Product Variety

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First version: April 2017. This version: August 2019

Abstract

This paper develops a new empirical method for valuing changes in product variety, what we define as the "variety effect". We show that the variety effect is pinned down by the causal effect of product variety on market demand. This is true in both a reduced-form model of demand that leaves preferences unspecified as well as a structural model of demand with generalized CES preferences. Since variety is endogenously determined, identifying the variety effect requires isolating exogenous variation in variety that is orthogonal to consumer preferences and other determinants of demand. We illustrate the method in an application based on consumer goods sold in grocery stores in the US. Central to our identification strategy are reduced-form effects of an instrument on price and quantity both when variety is held constant and when variety responds to the instrument, in addition to the effect of the instrument on variety. We also show how one may identify the variety effect in a two-stage least squares (2SLS) regression with the availability of two instruments, one for price and another for variety. Our instruments build on the recent literature on uniform pricing within retail chains. Across a wide range of specifications, our results consistently suggest that consumers place a large value on product variety.

JEL codes: H21, H71, F12, L13.

Keywords: Market demand, consumer surplus, product variety, new goods.

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1 Introduction

Quantifying the benefits to consumers from the introduction of new products is important for a broad range of economic issues, ranging from a full accounting of the gains from trade (Feenstra 1994; Broda and Weinstein 2006) to the welfare effects of tariffs (Romer 1994; Arkolakis et al. 2008), commodity taxes (Kroft et al. 2019a) and the socially optimal level of variety (Spence 1976a; Spence 1976b; Dixit and Stiglitz 1977; Mankiw and Whinston 1986; and Dhingra and Morrow 2019). The typical approach in international trade to measuring the "variety effect" – the gains from greater product variety, holding prices and cost constant – is to specify and estimate structural models of demand. Most analyses assume Constant Elasticity of Substitution (CES) preferences where the variety effect is determined by the elasticity of substitution combined with expenditure shares for the common varieties before and after the change in variety.[1] This framework has been very influential since it features nice aggregation properties and allows researchers to analyze the consequences of changes in product variety in a large number of markets simultaneously.

At the same time, it is well known that the CES framework has limitations. In particular, it links the taste for variety and market power (itself related to the degree of substitutability between products) together in an arbitrary way.[2] There is also a recognition that it may overstate the gains from variety.[3] Motivated by this, Benassy (1996) and Alessandria and Choi (2007) have considered more general models that disentangle the taste for variety from the elasticity of substitution and hence market power.[4] Despite progress on the theory front, there remains the empirical challenge of identifying the taste for variety as conceptually distinct from

[1]See Feenstra (1994); Broda and Weinstein (2006), Broda and Weinstein (2010); Arkolakis, Costinot and Rodriguez-Clare (2012); Handbury and Weinstein (2015); Melitz and Redding (2015); Atkin, Faber and Gonzalez-Navarro (2018).

[2]In the Dixit-Stiglitz-Spence model of monopolistic competition, this implies that the equilibrium level of production scale and the socially optimal level coincide as shown by Benassy (1996).

[3]For example, Feenstra and Weinstein (2017) consider translog preferences which account for crowding in the product space and find that the gains due to new product variety are overstated by a factor of 2 when considering CES preferences compared to translog preferences. Their results however indicate that the *overall* gains from trade (inclusive of pro-competitive effects and loss of variety due to exit of domestic firms) when considering translog preferences are similar to the overall gains when assuming CES preferences as in Broda and Weinstein (2006).

[4]We hereafter refer to the preferences in these models as "CES-Benassy preferences".

the degree of substitution across product varieties. The conventional approach of estimating the taste for variety using the price elasticity of demand is not valid in the presence of CES-Benassy preferences. Thus far, as Alessandria and Choi (2007) point out, there are no empirical estimates of the taste for variety which leads them to perform sensitivity analysis using different values of this parameter.

The main contribution of this paper is to address this empirical challenge. Theoretically, we show that the taste for variety can be identified using the *market demand for variety*. This is true both in a partial equilibrium model of demand with symmetric products which nests a wide range of preferences under a no income effects assumption as well as a structural model of demand with CES-Benassy preferences that accomodates product heterogeneity and income effects and thus permits a general equilibrium analysis. The first model uses aggregate demand for a set of differentiated products to derive the variety effect as the change in consumer surplus due to an exogenous change in the number of product varieties, holding market prices and the outside option fixed. The second model builds on Broda and Weinstein (2010) by adding Benassy preferences to the subutility function defined over the set of differentiated products and considers the exact price index (Feenstra 1994) which splits the change in welfare due to a change in variety into the variety effect which depends on the taste for variety and a price effect (the "conventional" exact price index, or CEPI)).

In both models, we show that the variety effect is fundamentally linked to the causal effect of variety on demand. Intuitively, if consumers value variety, then increases in variety for a set of goods should lead to greater demand for those goods, relative to the outside option. As such, the variety effect is pinned down by the amount that demand shifts in response to an exogenous change in variety. Whereas in a standard CES model, the causal effect of variety on quantity demanded and the causal effect of price on quantity demanded are effectively the same (and thus one can use the shortcut of identifying the variety effect using the price elasticity of demand), in a richer model with CES-Benassy preferences, they are conceptually distinct. This illustrates that identification of the variety effect requires separately identifying the causal effect of variety on quantity demanded independently from the effect of prices on quantity demanded.

We illustrate our method by considering an application to consumer products sold in grocery stores in the US. We focus on the partial equilibrium demand model which ignores income effects, for two reasons. First, the assumption of no income effects is likely a reasonable approximation in this setting, since most goods sold in grocery stores have small budget shares. Second, since we are the first to introduce this new method for valuing variety, it seems natural to start by focusing on this simpler setting as a "proof of concept". To operationalize our method, we recast the formula for the variety effect in terms of the causal effect of variety on willingness-to-pay (inverse aggregate demand). To link this formula to estimable parameters, we assume that the inverse market demand curve shifts *in parallel* in response to (exogenous) changes in variety relative to the (fixed) outside option.[5] Under parallel demands, we show that the variety effect can be identified using two sources of variation that are orthogonal to preferences and other determinants of demand, one source that only affects prices in the market and another source that affects both variety and prices. We use these insights to provide a graphical representation of the variety effect. As in Einav, Finkelstein, and Cullen (2010), we view these graphs as providing useful intuition and therefore as an important contribution on their own.

To generate variation in prices and variety that is orthogonal to consumer demand, we follow the recent literature on uniform pricing by retail chains (DellaVigna and Gentzkow 2019) and consider an instrument that is based on the pricing of products of other stores in a given retail chain. This instrumental variables strategy relates to the one in Hausman (1996) and Nevo (2001) and has been employed recently by Atkin, Faber and Gonzalez-Navarro (2018), DellaVigna and Gentzkow (2019) and Allcott et al. (2019). We show how one can use this pricing instrument both in the cross-section and the time-series to identify the variety effect. Intuitively, the pricing instrument in the cross-section affects both prices and variety. By contrast, we argue (and present empirical evidence) that this instrument

[5]The concept of "parallel demands" is developed in more detail in Kroft et al. (2109b), which shows that a wide class of discrete choice models of demand give rise to parallel demands.

mainly affects prices (and not variety) in the time-series. Under the assumption that the only difference between long-run demand and short-run demand is that variety is fixed in the short-run, estimating the effect of a supply-driven shock that only affects prices, holding variety constant, allows us to "back out" the hypothetical "ceteris paribus" variety shock from the shock that affects both prices and variety. Specifically, we show how one can estimate the variety effect using a "plug-in" estimator comprised of cross-sectional and time-series reduced-form estimates of the retail chain instrument on price and output. We also report a similar plug-in estimator based on a different set of reduced-form results from an altogether different source of exogenous variation in prices and variety: sales taxes. In particular, we use results in Kroft et al. (2019a) based on variation in sales tax rates rates across counties and sales tax exemptions within states to estimate the variety effect.[6] Both "plug-in" estimates suggest that consumers place a large value on product variety at grocery stores in the US.

To address the concern that the short-run price elasticity of demand may be smaller than the long-run price elasticity of demand for other reasons than endogenous product variety (e.g., other adjustment costs), we consider a complementary empirical approach. In particular, we also show how one can identify the variety effect using two-stage least squares (2SLS) and two instruments for price and variety. For price, we use the chain-level pricing instrument as in our first approach, and for variety we use a chain-level variety instrument which we show strongly predicts variety. Even though the pricing instrument affects variety, we generate independent variation in variety using the variety instrument and thus can identify both the causal effects of price and variety on demand.[7] The 2SLS estimates reveal a variety effect similar in magnitude to the plug-in estimates. Across all of our specifications, our results indicate a range of estimates for the variety effect of $0.7 - 1.2$, which corresponds to the

[6] Kroft et al. (2019a) estimates the effect of sales taxes on product variety and also uses the methodology developed in this paper to estimate the taste for variety using a different source of variation. Additionally, because taxes may not be fully salient, Kroft et al. (2019a) extend the methodology to show how to adjust the "plug-in" estimate developed in this paper for salience effects. Lastly, Kroft et al. (2019a) provide a new estimate of tax salience based on the instrumental variables estimates in this paper and the effects of sales taxes.

[7] While one can identify the parameters using these two instruments in a two-stage least squares regression, in practice, if the instruments are highly correlated, there may be a weak instruments problem.

elasiticity of average willingness-to-pay with respect to variety.

Although we relate our empirical estimates of the causal effect of variety on demand to the variety effect in the parallel demands model, in principle these estimates can be used to identify the variety effect in a broader class of models where the love of variety and price sensitivity are governed by two distinct parameters. Our derivation of the variety effect and the aggregate demand function in the CES-Benassy model is an attempt to illustrate this important conceptual point. In particular, we show that the aggregate demand function depends on a price index and variety. The coefficient on the price index is linked to the elasticity of substitution, while the coefficient on variety is linked to the variety effect. Thus, identification of the variety effect in this model can be achieved using the empirical strategy proposed in this paper -- by instrumenting for both price and variety and taking advantage of the separability in the aggregate demand function. From this standpoint, we see our contribution as coming up with a way to instrument for both price and variety and to demonstrate that the elasticities of demand with respect to price and variety are "sufficient" for welfare analysis in a broad class of models, including both models featuring parallel demands as well as models featuring CES-Benassy preferences.

The theoretical and empirical work in this paper relate to other topics in international trade. A central question in the trade literature is whether product creation is independent of market demand conditions as suggested by Schumpeter (1939) or whether it is driven by periods of high demand, as claimed by Schmookler (1962) and Shleifer (1986). Broda and Weinstein (2010) document a positive correlation between sales and product variety in the time-series and Hottman, Redding and Weinstein (2016) estimate a positive relationship between sales and the number of products under the assumption of CES preferences. Our contribution to this literature is to identify this relationship using reduced-form methods in a way that is model-free and in particular, does not rely on the CES assumption. Specifically, our uniform pricing instrument generates exogenous variation in demand which can then affect product variety. We find that the pricing instrument does in fact affect product variety, and by taking the ratio of two reduced-form estimates (the effect of the pricing instrument on variety and the effect of the pricing instrument on sales), we estimate the elasticity of product variety with respect to sales to be around $0.31 - 0.37$, which is very similar to the estimate in Broda and Weinstein (2010) of 0.35.

The remainder of the paper proceeds as follows. Section 2 considers models of demand and formally derives the variety effect. Section 3 considers identification of the variety effect. Section 4 considers an empirical application. Section 5 concludes.

2 The Variety Effect

In this section, we consider two approaches to identifying the welfare effects of product variety. The first approach relies on symmetry and partial equilibrium to use a revealed preference approach based on aggregate demand to identify the love of variety. The second approach drops symmetry and uses a more parametric approach based on Benassy (1996) and Alessandria and Choi (2007) to identify the variety effect in a general equilibrium model.

2.1 Parallel Demands

This section considers the partial equilibrium model with no income effects. All of the results we present rely only on demand and consumer surplus. The microfoundations for this model are in the Online Appendix. There is a single market with J varieties and an outside good and the price of product j in the inside market is given by p_j. We assume that there are no income effects and consider the case of symmetric preferences and a symmetric price equilibrium so that $p_j = p$ for all j. Under these assumptions, consumer surplus (up to a constant) is equal to the integral of theaggregate demand function$Q(p, J)$[8]:

$$CS(p, J) = \int_p^\infty Q(s, J)ds \tag{1}$$

[8]To see this, note that $CS(p_1, \dots, p_J, J) = \int_0^\infty \cdots \int_0^\infty Q(p_1 + r_1, p_2 + r_2, \dots, p_J + r_J, J)dr_1 dr_2 \cdots dr_J = \int_{p_J}^\infty \cdots \int_{p_1}^\infty Q(s_1, s_2, \dots, s_J, J)ds_1 ds_2 \cdots ds_J = \int_p^\infty \cdots \int_p^\infty Q(s, s, \dots, s, J)ds ds \cdots ds = \int_p^\infty Q(s, J)ds$ where the third equality uses a change of variables such that $s_i = p_i + r_i, ds_i = dr_i, \forall i = \{1, \dots, J\}$.

In what follows we assume that J is a continuous variable and $Q(p, J)$ is defined for any $J \in \mathbb{R}_+$ and is continuously differentiable.[9] We also denote the inverse aggregate demand corresponding to this function as $P(Q, J)$. We are interested in characterizing effects of a small increase in the number of varieties J on consumer surplus (holding the outside option constant). Applying Leibniz rule, the total effect of a small change in product variety on consumer surplus is:

$$\frac{dCS(p,J)}{dJ} = \frac{d}{dJ}\int_p^\infty Q(s,J)ds = \underbrace{-Q(p,J)\frac{dp}{dJ}}_{\text{price effect}} + \underbrace{\int_p^\infty \frac{\partial Q}{\partial J}(s,J)ds}_{\text{variety effect}} \tag{2}$$

Definition 1. The "price effect" is defined as:

$$-Q(p,J)\frac{dp}{dJ}. \tag{3}$$

Prices may change when varieties enter or exit the market. The envelope theorem implies that there is no first-order effect on utility due to consumer re-optimization when prices change, so only the mechanical effect of a price change affects consumer welfare.

Definition 2. The "variety effect" is defined as:

$$\Lambda(J) \equiv \int_p^\infty \frac{\partial Q}{\partial J}(s,J)ds. \tag{4}$$

Holding prices constant, an increase in variety increases consumer welfare since consumers exhibit a "love of variety". Equation (4) shows that the variety effect is related to the causal effect of product variety on aggregate demand. The next result shows that the variety effect (up to a first-order approximation) can be recast in terms of the causal effect of product variety on willingness-to-pay.

Lemma 1. *The variety effect can be equivalently represented as*

$$\Lambda(J) = \int_p^\infty \frac{\partial Q(s,J)}{\partial J}ds = \int_0^Q \frac{\partial P(t,J)}{\partial J}dt + O(dJ) \tag{5}$$

where $O(dJ)$ is a term which goes to zero at the same rate as dJ.

[9] In the Online Appendix, we extend the symmetric model to allow for probabilistic entry, where continuity of J is a more natural assumption. Additionally, if we allow for income effects, the correct measure of welfare is either compensating variation or equivalent variation. In that case, we can obtain an analogous result that compensating variation is the integral of the aggregate compensated demand.

Proof. See Online Appendix. □

The price effect and variety effect are illustrated in Figure 1 which considers a reduction in product variety in the market from J_0 to J_1 . The price effect is represented by the area $DFGH$ and the variety effect is given by the area $ABCD$. Intuitively, when the number of varieties is reduced, some consumers will no longer be able to purchase their most preferred option. Thus, the maximum willingness-to-pay for purchasing an inside good will be lower for these consumers. This is represented as a downward shift in the inverse aggregate demand curve. The area between the inverse aggregate demand curves before and after the change in variety (above the market price) corresponds exactly to the variety effect. Mapping this figure to equations (4) and (5), $\frac{\partial Q}{\partial J}$ represents the horizontal (quantity) distance between the two demand curves and $\frac{\partial P}{\partial J}$ represents the vertical (price) distance. Integrating the quantity distance (area $ABCD$) and integrating the price distance (area $ABED$) lead to the same measure that differs up to the small area labelled CDE, which corresponds to the term $O(dJ)$ in equation (5). The next result shows that this term is approximately zero when the change in variety is small.

Corollary 1. *Consider a small change in the number of varieties. In this case, the variety effect can be represented in terms of the average change in willingness-to-pay for inframarginal units as:*

$$\Lambda(J) \approx \int_0^Q \frac{\partial P(t,J)}{\partial J} dt = Q \overline{\frac{\partial P}{\partial J}}(Q,J), \tag{6}$$

where the average change in willingness-to-pay for inframarginal units is

$$\overline{\frac{\partial P}{\partial J}}(Q,J) \equiv \frac{1}{Q} \int_0^Q \frac{\partial P}{\partial J}(t,J) dt.$$

Proof. See Online Appendix. The key idea is to use a Taylor expansion to expand the bias term and approximate this term with first-order and second-order changes in J. □

A key objective of this paper is to establish a method to identify the variety effect using reduced-form methods. To make progress on this, we introduce the following assumption.

Assumption 1. *The inverse aggregate demand curves are parallel, such that* $\frac{\partial P}{\partial Q}(Q, J) = \frac{\partial P}{\partial Q}(Q, J')$ *for all* J, J' *and* Q.

Assumption 1 implies that the average change in willingness-to-pay is equal to the marginal change in willingness-to-pay, $\overline{\frac{\partial P}{\partial J}}(Q, J) = \frac{\partial P}{\partial J}(Q, J)$ for all Q and J. Kroft et al. (2019b) show that a large class of conventional discrete choice models of demand are good approximations to Assumption 1. The next theorem states that one may recover the variety effect using a revealed preference approach under this assumption.

Theorem 1. *If the change in the number of varieties is small and Assumption 1 holds, then*

$$\overline{\frac{\partial P}{\partial J}} = \frac{\partial P}{\partial J} = \frac{dP}{dJ} - \frac{\partial P(Q, J)}{\partial Q}\frac{dQ}{dJ} = \frac{dQ}{dJ}\left(\frac{\frac{dP}{dJ}}{\frac{dQ}{dJ}} - \frac{\partial P(Q, J)}{\partial Q}\right) \tag{7}$$

for all Q *and* J*, where* $\frac{\partial P(Q,J)}{\partial Q}$ *denotes the slope of inverse demand when variety* J *is held fixed and* $\frac{dP}{dJ}/\frac{dQ}{dJ}$ *denotes the slope of inverse demand when* J *is variable.*

Proof: See Online Appendix.

The expression for the variety effect in equation (6) and the average change in willingness-to-pay in equation (7) can be most easily understood geometrically. Figure 2 considers a reduction in variety from J_0 to J_1 and shows that the variety effect is the shaded area of the rectangle with base Q_0 and height $d \equiv P_1 - P^*$. The base Q_0 is aggregate output prior to the change in variety. The height d captures the change in willingness-to-pay as J changes. Note that d is not directly observable since it depends on the market price that would prevail at the final level of output but on the original demand curve, defined as P^*, which is unobservable. To see how to recover an expression for d, note from Figure 2 that d must satisfy the following relationship $Q(P(J_1), J_1) = Q(P(J_1) - d, J_0)$. Let $P_1 \equiv P(J_1)$. We can solve for d as follows:

$$\begin{aligned}
dQ &= Q(P_1, J_1) - Q(P_0, J_0) \\
dQ &= Q(P_1 - d, J_0) - Q(P_0, J_0) \\
dQ &\approx \frac{\partial Q(P, J)}{\partial P}|_{P=P_0}(-d + P_1 - P_0) \\
dQ &= \frac{\partial Q}{\partial P}|_{P=P_0}(-d + \frac{dP}{dQ}dQ)
\end{aligned}$$

The first equality holds by definition. The second equality holds by Assumption 1. The third approximation holds by doing a Taylor expansion of $Q(P, J)$ around P_0. The fourth equality holds by definition. Rearranging and solving for d yields:

$$d \approx dP - \frac{\partial P}{\partial Q} dQ$$

In economic terms, d gives the reduction in the willingness-to-pay for the marginal unit. Under Assumption 1, it also gives the change in willingness-to-pay for inframarginal units. In order to empirically implement this expression, two partial effects are required. First, one needs an estimate of demand sensitivity, holding variety fixed, $\frac{\partial P}{\partial Q}$. Second, one requires reduced-form estimates of the change in price (dP) and output (dQ) in response to a change in variety. While these partial effects are local since they are evaluated at the equilibrium, one still needs to extrapolate in order to recover the willingness-to-pay for the infra-marginal consumers. Thus, the revealed preference approach is not strictly speaking local, but rather uses local estimates to extrapolate globally. The upshot is that it's possible to identify d using the price elasticities of demand when J is fixed and when J is variable. This requires two supply-side sources of variation, one which varies prices holding variety constant, and another varying variety (which may also impact prices), as we describe more fully in section 3. Finally, although we considered the case of symmetric products, the results of the section apply more broadly. In the Online Appendix, we consider the case of asymmetric products and show that under the assumption that all prices adjust uniformly after the introduction of the new varieties, one can still identify the variety effect using equation (6).

2.2 CES-Benassy Preferences

A limitation of the parallel demands approach to quantifying the gains from variety is the assumption of symmetric products and no income effects. In our empirical application which focuses on product variety in grocery stores, assuming no income effects may be a reasonable approximation, but this approach may be potentially less suitable for a general equilibrium analysis which is the focus of the trade literature. This subsection considers a model with

CES-Benassy preferences that builds on the framework in Broda and Weinstein (2010) and has the advantage of permitting a general equilibrium analysis where preferences might not be symmetric and changes in variety in a single market may affect other markets. The primary objective is to derive the variety effect in a model more familiar to the trade literature and to show how one can identify the variety effect using similar sources of variation to those described in the last section; in particular, one source that only affects prices in the market and another source that affects both prices and variety.

The upper level utility function takes the following form:

$$U = (q_o^{\rho_1} + Q^{\rho_1})^{\frac{1}{\rho_1}} \tag{8}$$

where q_o is the outside good and $\sigma_1 \equiv \frac{\rho_1}{\rho_1 - 1} < 0$ is the elasticity of substitution across products groups (the inside goods and the outside good).[10] We model the lower tier as follows:

$$Q = J^{\nu+1-\frac{1}{\rho_2}} \left(\sum_{j=1}^{J} (b_j q_j)^{\rho_2} \right)^{\frac{1}{\rho_2}} \tag{9}$$

Note that if $\nu = \frac{1}{\rho_2} - 1 = -\frac{1}{\sigma_2}$, we get back to standard CES preferences with an outside good. This parameter represents the taste for variety (Benassy 1996 and Alessandria and Choi 2007). The parameters $\{b_j\}_{j=1}^{J}$ capture unobserved product quality. Income is given exogenously as I and we normalize the price of the outside good to $p_o = 1$.

In the standard nested-CES model following Broda and Weinstein (2010) where $\nu = \frac{1}{\rho_2} - 1$ the minimum unit cost function associated with the subutility function in equation (9) is given by:

$$P_J^{CES} = \left(\sum_{j=1}^{J} \left(\frac{p_j}{b_j} \right)^{\sigma_2} \right)^{\frac{1}{\sigma_2}} \tag{10}$$

In the CES-Benassy model, we have:

$$P_J = \frac{1}{J^{\nu+\frac{1}{\sigma_2}}} P_J^{CES} \tag{11}$$

In the case where products are symmetric ($b_j = 1$) and varieties are equally priced ($p_j = p$),

[10] In the trade literature, the elasticity of substitution is usually defined as either $\sigma = \frac{\rho}{\rho-1}$ or $\sigma = \frac{1}{1-\rho}$. Here, we adopt the first definition. Additionally, in principle we could include other nests of differentiated products, as in Broda and Weinstein (2010), but we focus on a single nest for tractability and for ease of exposition.

the minimum unit-cost functions become $P_J^{CES} = J^{\frac{1}{\sigma_2}}p$ and $P_J = \frac{1}{J^{\nu}}p$. As J increases, the minimum cost required to attain a given level of utility falls. In the case where $\nu = -\frac{1}{\sigma_2}$, the welfare gain is pinned down by the elasticity of substitution σ_2.

The cost function associated with the outside good is $P_o = 1$. The overall price index associated with equation (8) is thus given by:

$$P = (1 + P_J^{\sigma_1})^{\frac{1}{\sigma_1}} \tag{12}$$

2.2.1 Deriving the Variety Effect

The typical approach in trade to evaluating the welfare effect of varieties when preferences are of the CES form is to use the ideal price index which is the ratio of two expenditure functions (Feenstra 1994).[11] This is because under homothetic preferences, as is the case with CES, the ratio of two expenditure functions does not depend on utility. By contrast, equivalent variation does depend on utility in this case and hence is not a convenient measure of welfare.[12]

We follow the trade approach and define the Exact Price Index (EPI) of the overall market allowing for entry and exit of goods in the inside market and assuming that tastes or product quality and the price and quality of the outside good do not change between periods $t-1$ and t:

$$EPI\left(\mathbf{p}_t, \mathbf{p}_{t-1}, J_t, J_{t-1}\right) = \left(\left(\frac{J_{t-1}}{J_t}\right)^{\nu+\frac{1}{\sigma_2}} CEPI\left(\frac{s_t^{Com}}{s_{t-1}^{Com}}\right)^{-\frac{1}{\sigma_2}}\right)^{w_J}$$

where $\overline{J} = J_t \cap J_{t-1}$, $CEPI = \prod_{j\in\overline{J}}\left(\frac{p_{j,t}}{p_{j,t-1}}\right)^{w_j}$ with $w_j = \frac{\frac{\overline{s}_{j,t}-\overline{s}_{j,t-1}}{log\overline{s}_{j,t}-log\overline{s}_{j,t-1}}}{\sum_{j\in\overline{J}}\frac{\overline{s}_{j,t}-\overline{s}_{j,t-1}}{log\overline{s}_{j,t}-log\overline{s}_{j,t-1}}}$, $s_r^{Com} = \frac{\sum_{j\in\overline{J}} p_{j,r}q_{j,r}}{\sum_{j\in J_r} p_{j,r}q_{j,r}}$, $r = \{t-1, t\}$ and $w_J = \frac{\frac{s_{J,t}-s_{J,t-1}}{logs_{J,t}-logs_{J,t-1}}}{\frac{s_{o,t}-s_{o,t-1}}{logs_{o,t}-logs_{o,t-1}}+\frac{s_{J,t}-s_{J,t-1}}{logs_{J,t}-logs_{J,t-1}}}$. The weights w_j and w_J are the log-ideal CES Sato (1976) and Vartia (1976) weights. As in Broda and Weinstein (2010), the result indicates that the exact price index can be decomposed into a term CEPI (the "conventional"

[11] As we shall see below, this combines both the price effect and the variety effect.

[12] Note that in section 2.1, the variety effect Λ is based on consumer surplus which is also equal to equivalent variation under the assumption of no income effects. This means that the measures of the variety effect across models will not be entirely consistent, although they both measure the welfare gains of new varieties, holding prices constant.

exact price index) which captures the price changes of goods that are common to both periods and a share ratio which captures the effect of exit and entry on the shares of common goods in the market. The new term which depends on the ratio of varieties reflects the impact of Benassy preferences and the taste for variety. Taking logs of the EPI, we may decompose it as follows:

$$logEPI = w_J logCEPI + w_J \left(\nu + \frac{1}{\sigma_2}\right) log\left(\frac{J_{t-1}}{J_t}\right) - w_J \frac{1}{\sigma_2} log\left(\frac{s_t^{Com}}{s_{t-1}^{Com}}\right)$$

This equation illustrates an advantage of the structural trade approach, namely that it permits heterogeneity across products. To the extent that there are unobserved quality differences across products, these will be captured by expenditure shares. The simplest case is when products are symmetric. Under symmetry, the EPI can be written as:

$$\frac{logEPI}{logJ_{t-1} - logJ_t} = \underbrace{w_J \frac{logCEPI}{logJ_{t-1} - logJ_t}}_{price\,effect} + \underbrace{w_J \nu}_{variety\,effect} \tag{13}$$

The first term in equation (13) represents the price effect and the second term represents the variety effect.[13] One can see that the change in welfare due to the change in variety relative to the (log) change in variety– holding prices constant – is pinned down by $w_J v$. The weight w_J in practice is measurable using data on expenditure shares. Thus, to identify the variety effect, one has to recover the taste for variety parameter ν.

2.2.2 Identifying the Variety Effect

It is useful to contrast this with the conventional CES case where $\nu = \frac{1}{\rho_2} - 1 = -\frac{1}{\sigma_2}$. Here one can see that the variety effect is pinned down by the degree of substitutability between the goods. The standard approach to identifying the variety effect in this case is to estimate the price elasticity of demand. To see this, consider the demand structure implied by the model. The expenditure on product j is given by:

$$p_j q_j = \left(\frac{P_J}{P}\right)^{\sigma_1} \left(\frac{\frac{p_j}{b_j}}{P_J^{CES}}\right)^{\sigma_2} I$$

[13] The second term represents the variety effect for a discrete change in J. This contrasts with the prior section which presented the variety effect for a marginal change in J.

The expenditure share on product j in the inside market is given by:

$$s_j = \frac{p_j q_j}{\sum_{j=1}^{J} p_j q_j} = \left(\frac{\frac{p_j}{b_j}}{P_J} \right)^{\sigma_2}$$

Taking logs, we get:

$$log s_j = \sigma_2 \left(log \frac{p_j}{b_j} - log P_J \right) \tag{14}$$

Therefore, equation (14) shows that one can recover the taste for variety using exogenous price variation *at the product level*. For example, Atkin, Faber and Gonzalez-Navarro (2018) use an instrument for product-level price and include market fixed effects to sweep out P_J in order to identify the elasticity of substitution.

In the more general CES-Benassy case, the price elasticity of demand is no longer sufficient to compute the taste for variety. We argue that the aggregate expenditure share on products in the inside market should be used instead. This takes the following form:

$$s_J = \sum_{j=1}^{J} s_j = \left(\frac{P_J}{P} \right)^{\sigma_1}$$

Taking logarithms of the aggregate share, we get:

$$log s_J = \sigma_1 log P_J / P$$

Substituting using equation (11) yields:

$$log s_J = -\sigma_1 \left(\nu + \frac{1}{\sigma_2} \right) log J + \sigma_1 \left(log P_J^{CES} - log P \right) \tag{15}$$

Equation (15) shows that demand depends on prices in the market relative to the overall price index in addition to variety. As prices in the inside market increase relative to the price of the outside option, demand falls, and the magnitude of the shift is determined by the elasticity of substitution, σ_1. Equation (15) shows that the coefficient on log variety is jointly pinned down by the parameters $(\sigma_1, \sigma_2, \nu)$. As described above, σ_2 can be identified using the price elasticity of demand at the product level. Next, σ_1 and ν may be recovered as the causal effects of average prices and variety in the market, respectively. In order to identify these parameters, one requires instrumental variables for prices and variety. Intuitively, it makes sense that the aggregate demand function is required to identify the taste for variety since this term operates at the nest level in the utility function in equation (9). In practice, it may be difficult to find an instrument that "only" shifts variety and not price; as a result, in the next section we describe an identification approach that can be used both for the parallel demands approach and the CES-Benassy model developed in this subsection.

3 Identification and Estimation of the Variety Effect

This section describes our approach to identifying and estimating the variety effect. For convenience, we focus throughout on the variety effect defined in subsection 2.1. However, the key empirical moments that we introduce can also be used to identify the variety effect in the CES-Benassy model. Intuitively, our identification approach relies on the availability of an instrumental variable observed in two different scenarios: (1) the instrument affects prices but not product variety; and (2) the instrument affects both prices and variety. In both scenarios, we require that the instrument affects the prices for the products in the "module" (or product market), and the identifying assumption is that the variation in prices and variety arises from supply-side variation, and is thus orthogonal to preferences and demand. In the first scenario, the instrument generates variation in prices but not variety (we will call this "short-run" variation, since varieties may not adjust immediately in response to ashock that affects prices); in the second scenario, the instrument can either generate variation in prices and/or variety directly, or indirectly through firm responses to the instrument. We will call this "long-run" variation since it allows for firm responses. For example, consider a shock to the fixed cost of entry of a new variety. This will directly affect variety, and in turn affect equilibrium prices (depending on the specific model of supply-side competition). This would be a valid instrument for the second scenario. Equation (15) shows that these two exogenous sources of variation are exactly what is required to estimate the taste for variety, ν.[14]

Formally, we consider an instrument z that is observed in two scenarios: (1) when variety is held fixed and (2) when variety can vary in order to identify the variety effect as follows:

$$\Lambda = Q\left[\frac{\frac{dp}{dz}}{\frac{dQ}{dz}} - \frac{\left.\frac{dp}{dz}\right|_J}{\left.\frac{dQ_J}{dz}\right|_J}\right]\frac{\frac{dQ}{dz}}{\frac{dJ}{dz}}. \tag{16}$$

To see where the above expression comes from, note that $\Lambda \approx d{*}Q$ where $d = \left(\frac{dP}{dQ} - \left.\frac{dP}{dQ_J}\right|_J\right) dQ$. Intuitively, equation (16) comes from replacing the dP/dQ terms with Wald estimators using theinstrument z. Since the instrumental variable is a supply-side instrument, any instrument that generates exogenous variation in prices and holds variety fixed will trace out the fixed-variety demand curve; thus, the ratio of $\left.\frac{dp}{dz}\right|_J$ to $\left.\frac{dQ_J}{dz}\right|_J$ will be a valid estimate of $\left.\frac{dP}{dQ_J}\right|_J$. For the dP/dQ term, we replace it with the ratio of dp/dz to dQ/dz. An instrumental variable that affects prices and/or variety directly – and in turn generates endogenous supply-side price and variety responses to the initial shock from the instrument – will lead to valid reduced-form estimates dp/dz and dQ/dz which can be used to calculate the dP/dQ term. Recall that the conceptual framework in subsection 2.1 defines Λ as the willingness-to-pay for an exogenous change in variety *holding prices constant*; this formula makes clear that it is still possible to identify Λ even if that exact conceptual experiment is not available and instead there is an instrument that affects both prices and variety. The key to the expression above is the ability to "subtract out" the component of the reduced-form estimate that comes only from prices (holding variety constant). This is the essence of our identification approach.

The intuition for the variety effect can be illustrated by returning to Figure 2 and re-writing the formula in equation (16) as follows:

$$\Lambda = Q * \left[\frac{dp}{dz} - \left.\frac{dp}{dz}\right|_J \frac{\frac{dQ}{dz}}{\left.\frac{dQ}{dz}\right|_J}\right]\frac{1}{\frac{dJ}{dz}}$$

In Figure 2, the base of the rectangular area is given by the pre-existing output before the

[14]Its possible that the short-run and long-run demands are different for reasons that have nothing to do with variety, such as adjustment costs. In the Online Appendix, we extend the model in 2.1 to consider the case where there is an outside market represented by a variable y. We use the principle of Le Chatelier to show that under the assumption that consumers can only adjust y in responses to changes in variety in the long-run, but not the short-run, our formula for the variety effect remains valid.

change dz (Q). The height of the rectangle is given by the difference between the "long-run" change in price and the "short-run" change in price re-scaled by the ratio of the long-run output effect to the short-run output effect of the instrument (i.e., $dp/dz|J$ scaled by $\frac{dQ}{dz} / \left.\frac{dQ}{dz}\right|_J$). The re-scaling serves to extend the price effect up the demand curve so that it's measured at the same long-run output level, and dividing by dJ/dz re-scales the entire expression in order to define the variety effect to be for a given change in variety. This discussion emphasizes that the identification of the variety effect comes from an instrumental variable (such as tariffs or sales taxes) that shifts both prices and varieties in the market, and requires both a setting where variety is held constant and a setting where variety is responsive to the instrument.

This discussion highlights a key strength of our framework: equation (16) shows how to recover variety effect – even in situations when the shock (dz) affects *both* prices and varieties in a market. In other words, our framework does not require the existence of an instrument which *only* affects variety in a market (i.e., shifts variety while holding prices constant) which might be difficult to find in practice. Our approach instead shows how to combine "fixed-variety" reduced-form estimates with "long-run" reduced-form estimates, where variety responds endogenously (in response to shocks that change producer profits).

To provide some intuition for the variety effect, Table 1 provides a set of illustrative calibrations. Across all columns, we report the variety effect as an elasticity, defined as $\tilde{\Lambda} \equiv \frac{J}{pQ}\Lambda$ and as follows:

$$\widetilde{\Lambda} \equiv \frac{J}{pQ}\Lambda = -\left[\left.\frac{d\log(p)}{dz}\right|_J \frac{\frac{d\log(Q)}{dz}}{\left.\frac{d\log(Q)}{dz}\right|_J} - \frac{d\log(p)}{dz} \right] \frac{1}{\frac{d\log(J)}{dz}} \tag{17}$$

From now on, we will focus on $\widetilde{\Lambda}$, which corresponds to the elasticity of consumer surplus (as a share of revenue) with respect to variety. The scaling of the variety effect gives a unit-free measure that can be interpreted as the elasticity of the inverse aggregate demand curve with respect to variety. This has the advantage of being easily implementable as a "plug-in" estimator, since our main reduced-form estimates will be based on regressions that take logs of prices, quantity, and variety.

Table 1 reports different estimates of $\tilde{\Lambda}$ for different combinations of reduced-form estimates. In column (1), the short-run and long-run effects of the instrument on prices and output are the same, and so the variety effect is 0 (and this would be true regardless of the estimated value of $d\log(J)/dz$). In this case, the reduced-form estimate reveals that consumers do not value variety. Since the shift in the aggregate demand curve depends jointly on the variety response to the instrument ($d\log(J)/dz$) and the variety effect ($\tilde{\Lambda}$), then if $\tilde{\Lambda} = 0$ there is no downward shift in the aggregate demand curve, despite the fact that the instrument nevertheless affects varieties (i.e., $d\log(J)/dz < 0$). Turning to column (2), in this scenario the long-run output elasticity exceeds the short-run output elasticity and this leads to an implied estimate of the variety effect of 1. Column (3) uses the same price and output estimates as column (2) but considers a smaller (in magnitude) estimate for $d\log(J)/dz$; since all of the price and quantity estimates are the same but the reduced-form effect on variety is smaller, this translates into a larger estimate of $\tilde{\Lambda}$. Intuitively, for a smaller change in variety to be associated with the same "difference" between thelong-run and short-run demand elasticities, it must be the case that consumers care more about variety, because otherwise the aggregate demand curve would not have shifted out as much in response to the change in variety. Lastly, column (4) shows a combination that leads to a smaller estimated variety effect. This is a smaller variety effect than column (2) because it keeps all reduced-form estimates the same except for the long-run estimate, which is reduced from -1.6 to -1.3. This leads to a smaller estimated variety effect, which corresponds graphically to a smaller inward shift in demand and thus a smaller area between the two inverse aggregate demand curves.

4 Empirical Implementation

The previous section showed that the causal effects needed to identify the variety effect are $\frac{d\log(J)}{dz}$, $\frac{d\log(p)}{dz}$, $\frac{d\log(p)}{dz}\Big|_J$, $\frac{d\log(Q)}{dz}$, and $\frac{d\log(Q)}{dz}\Big|_J$. This section presents empirical estimates of these effects. Our goal is estimate the variety effect for a set of differentiated products in grocery stores in the U.S. For measures of p, Q, and J, we use the Nielsen Retail Scanner

Data, which records weekly prices and sales by product (i.e., by Universal Product Code (UPC)) for stores across the U.S. from $2006 - 2014$. Each UPC in the data belongs to a product "module", which corresponds to a broad category of products (e.g., breakfast bars, olive oil).[15]

Starting from a sample of $11,487$ grocery stores, we impose several sample restrictions. First, we only include stores that are assigned to the same retail chain throughout the $2006 - 2014$ period, that are present in the data for at least two years, and that belong to retail chains that were associated with the same parent company throughout the period. Second, out of the $1,102$ modules present in Nielsen's Retail Scanner Data, we focus on the top 20% modules in terms of total sales. These modules account for almost 80% of the total value of sales in grocery stores in the scanner data. Our final sample includes 198 modules sold in $8,653$ grocery stores, from 70 retail chains, over 9 years (36 quarters).

We aggregate the UPC-level micro data to the store-module-time level to create two samples, one for the cross-sectional analysis (frequency = yearly) and another for the time-series analysis (frequency = quarterly). Price is constructed as a time-varying module-store level index. To account for heterogeneity in product quality, we estimate regressions of log prices on UPC and store-by-time fixed effects using the micro data, separately for each module (as in Handbury and Weinstein 2015). We recover the store-by-time fixed effects and use them as adjusted store-module measures of price ($\log p_{mrt}$), where m = module, r = store, and t = time (yearly for the cross-sectional analysis, and quarterly for the time-series analysis). To define output Q_{mrt}, we aggregate UPC-level expenditures within store-module-time cells fixing each product's price at its national average. Since prices are held constant, variation in Q_{mrt} across stores and over time reflects variation in quantities.[16] Finally, variety J_{mrt} is

[15]See Table A1 in the Appendix for examples of UPCs and the organizational hierarchy of the Nielsen data.

[16]This approach is similar to the real consumption index developed by Kaplan, Mitman and Violante (2016). Were we to use UPC-level data in our analyses, instead of module-level data, the output elasticities obtained using Q_{jmrt} (where j = UPC) as our dependent variable would be numerically identical to elasticities estimated using the actual quantities sold q_{jmrt} because of the inclusion of product (UPC) fixed effects. Alternatively, we could focus on log expenditure (i.e., log of $p_{jmrt} * Q_{jmrt}$) instead of our preferred quantity measure, and then use the chain rule to "back out" the effect of instrument on quantity from the effect of instrument on log expenditure and log average price (i.e., $\frac{dlogQ}{dz} = \frac{dlogpQ}{dz} - \frac{dlogp}{dz}$). We have done this analysis and the results are very similar for all of our main results.

defined as the count of UPCs with positive sales within a module and store over the relevant time period.[17]

Key to our analysis is an instrument for price, z. Our main instrument builds on DellaVigna and Gentzkow (2019) who demonstrate that for a large share of UPCs sold in grocery stores, at any point in time, prices are set uniformly across stores within retail chains. Formally, for each observation in our data (at the store-module-time level), we construct an instrument that is equal to the average log price (within each module m) for all stores in the same chain excluding store r:

$$\log p_{mft,-r} = \frac{\sum_{s\in f} \log p_{mst} - \log p_{mrt}}{N_f - 1}$$

where f denotes the retail chain to which store r belongs and N_f is the number of stores in chain f. This is a valid instrument under the assumption that chain-specific prices predict store prices, but are not correlated with unobserved store-level demand. A threat to the validity of our research design is there may be correlated demand shocks across stores within chains. To address this, we include store fixed effects in all of our specifications. We also include module fixed effects to account for the fact that more expensive modules may reflect chains responding to strong demand for these modules. Thus, our identification is coming from differences in relative prices across modules and across chains. To the extent that this variation is driven by differences in product-specific marginal costs across chains, differences in distribution costs across chains (such as supply-sourcing costs), or differences in bargaining power across chains, we can consistently estimate our reduced-form effects of interest (DellaVigna and Gentzkow 2019). One key identifying assumption is that chains select store locations based on overall demand (across modules), but not module-specific demand.

While our main analysis uses the "leave-me-out" average log price instrument defined above, we will also briefly summarize results from an alternative instrumental variable which uses sales taxes to instrument for prices and varieties. This analysis re-uses the empirical results in Kroft et al. (2019a), which describes the sales tax analysis in much more detail.

[17]More details regarding the construction of each variable are provided in the Data Appendix.

Essentially, we replicate our main results using sales taxes instead of leave-me-out average prices and estimate the same reduced-form statistics needed to implement the variety effect estimator in equation (16).

Lastly, we introduce a new instrumental variable also building on DellaVigna and Gentzkow (2019); in that paper, the authors show that chains set prices uniformly and also – but to a somewhat lesser extent – chains set varieties uniformly.[18] As a result, we construct an instrument for variety that is analogous to the average log price instrument:

$$\log J_{mft,-r} = \frac{\sum_{s \in f} \log J_{mst} - \log J_{mrt}}{N_f - 1}$$

We use this instrument along with the average log price instrument in a two-stage least squares (2SLS) analysis below as an alternative way of identifying the variety effect. Intuitively, the leave-me-out average log price instrument "primarily" affects prices (but may also indirectly affect variety), while the leave-me-out variety instrument "primarily" affects variety. Thus, these two instrumental variables can be used as instruments for prices and variety; we describe the connection to our main identification approach below.

4.1 Cross-Sectional Estimates of $\frac{dlog(p)}{dz}$, $\frac{dlog(Q)}{dz}$, and $\frac{dlog(J)}{dz}$

We estimate the cross-sectional reduced-form effects of interest using the following regression model:

$$\log y_{mrf} = \beta^{LR} \log p_{mf,-r} + \delta_r + \delta_m + \varepsilon_{mrf} \tag{18}$$

[18]We verify this pattern in our own analysis of the Nielsen data: we find that product variety is less uniform across stores within chains than prices are. For instance, in Appendix Table A7 we show that while stores' average prices are five-time more responsive to the average income of consumers of their chain than to their own consumers' income, the corresponding ratio is much smaller for product variety. In a store-level regression of product variety on consumer income, the coefficient on chain-level average income is not statistically significant (0.244, s.e. 0.206) and is less than twice the size of the coefficient on own-store income (0.137, s.e. 0.039). Similarly, we find that about 80% of residual variation in prices is between-chain, whereas only 63% of the residual variation in variety is between-chain. We calculate this based on residuals from regressions of store-module yearly measures of log prices (log variety) on module-by-year fixed effects. We then calculate the fraction of the total variance in these residuals that occur within chain-by-module-by-year cells.

where the outcome y_{mrf} is either price (p_{mrf}), quantity (Q_{mrf}), or variety (J_{mrf}), and δ_r and δ_m are store and module fixed effects, respectively. The store-by-module-level instrument, $\log p_{mf,-r}$ is the average log price of module m, for all stores of retail chain f, except for store r. Thus, $\log p_{mf,-r}$ serves as instrumental variable (z), and each of the β^{LR} coefficient estimates corresponds to one of the three reduced-form effects needed to implement the variety effect formula. We estimate equation (18) separately for each yearly cross-section between 2006 and 2014, and we then take a simple linear combination of all the coefficient estimates, putting equal weight on all 9 yearly cross-sectional estimates. Standard errors are clustered at the chain-module level.[19]

We report the main estimates in Table 2. The dependent variable is log average price in columns (1) and (2), log quantity in columns (3) and (4), and log variety in columns (5) and (6). Columns (1), (3) and (5) present results for specifications which control for module fixed effects and store fixed effects; module-by-county fixed effects are further added in columns (2), (4) and (6) to address the possibility that stores belonging to high-price chains may locate in particular markets in a way that is correlated with product-specific local demand characteristics. The results in column (1) shows that the elasticity of average log price with respect to the chain instrument is very close to 1 and is very precisely estimated; this verifies that the instrument has an extremely strong first stage, as one would expect given existing work on uniform pricing within chains (DellaVigna and Gentzkow 2019). The estimate falls somewhat with additional fixed effects, as shown in column (2), but remains very precisely estimated. Column (3) shows that the elasticity of quantity demanded with respect to the chain instrument is roughly -1.9 (s.e. 0.12) in the main specification, and column (4) shows that this estimate falls to -1.5 (s.e. 0.08) with the additional fixed effects, and the implied price elasticity of demand is similar across the two columns (-1.9 and -1.7). Lastly, column (5) shows that the elasticity of variety with respect to the chain instrument is -0.69 (s.e. 0.08), and this estimate falls somewhat to -0.48 (s.e. 0.05) with additional fixed effects.

[19]Econometrically, we obtain the chain-module clustered standard errors for the equally-weighted linear combination of the yearly cross-sectional estimates using a single estimating equation that simultaneaously estimates our 9 cross-sectional reduced-form effects.

In Appendix Table $A2$, we report a variety of robustness checks for these main results. To verify that our results do not hinge on using counties to operationalize local markets (when including module-by-county fixed effects), we allow the module fixed effects to vary flexibly across markets using an alternative definition (Nielsen's Designated Market Areas, or DMAs). For all three dependent variables, we find very similar elasticities. Another potential threat to the validity of our results is that stores of a given chain may tend to locate in the same market, so that variation in our instrument may still reflect variation in market-level demand (which would not be captured by market-specific module effects). To ensure that this is not driving our results, we construct alternative "uniform pricing" instruments using only stores in the same chain located in other markets (i.e., in other DMAs or in other counties). Our estimates are virtually unchanged using these alternative instruments, indicating that the reported elasticities are plausibly reflecting differences in relative costs, not differences in relative demand in local market. Importantly, including store and module fixed effects throughout our analysis is key to the high amount of robustness to these additional controls. Without store fixed effects, we find that our main results are much more sensitive to these controls, suggesting that the reliability of results comes from exploiting variation across modules within a store. We also explore an alternative measure of product variety in Appendix Table $A3$. In our main results we focus on the "raw" number of distinct UPC codes as our preferred measure of product variety. However, marginal varieties might have systematically different market shares. We evaluate this empirically by calculating a "share-weighted" measure of variety that imputes market shares for each UPC code in a given store and module combination using market shares for that same UPC code in other markets. Using this alternative variable, we continue to find a negative effect of average log prices on variety, but the effect is somewhat smaller in magnitude, suggesting that marginal varieties have lower-than-average market shares.

While our primary focus is using these reduced-form estimates to measure the variety effect, our estimates can also be used to connect to the work of Broada and Weinstein (2010) and Hottman, Redding and Weinstein (2019), which finds a positive relationship between sales and the number of products (in the latter case, under the assumption of CES preferences). Our estimates can be used to provide a model-free estimate of this relationship, since our uniform pricing instrument generates exogenous variation in demand which can then affect product variety. We find that the pricing instrument does in fact affect product variety, and by taking the ratio of two reduced-form estimates, we find that the elasticity of product variety with respect to sales to be around 0.31-0.37, which is very similar to the estimate in Broda and Weinstein (2010) of 0.35.

4.2 Fixed-Variety Estimates of $\frac{d\log(p)}{dz}\Big|_J$ and $\frac{d\log(Q)}{dz}\Big|_J$

To estimate the reduced-form effects of interest, we need to generate variation in prices holding variety fixed. We estimate these short-run reduced-form effects using the panel structure of data covering the $2006-2014$ time period and estimating regression models that control for module fixed effects interacted with store fixed effects (so that the estimates are based on variation over time within a store and module combination). Formally, the econometric specification is the following:

$$\log y_{mrft} = \beta^{SR} \log p_{mft,-r} + \delta_{mr} + [\delta_{mt} + \delta_{rt}] + \varepsilon_{mrft} \tag{19}$$

where $\log p_{mft,-r}$ is the leave-me-out chain average log price, but in this case measured separately each quarter. With the inclusion of module-by-store fixed effects (δ_{mr}), we are effectively identifying β^{SR} from comparing changes in prices and quantity over time for a given module within a given store. We further add module-by-quarter fixed effects (δ_{mt}) to allow product categories to have arbitrary time trends in some specifications, and in some specifications we also add store-by-quarter fixed effects (δ_{rt}) to account for any store-specific (or chain-specific) shocks that affect all modules simultaneously.

The main results are reported in Table 3, and the layout follows Table 2: columns (1) to (2) contain our estimates for prices and columns (3) to (4) contain our results for quantity. In odd-numbrred columns we include module-by-quarter fixed effects, and in even-numbered columns we add in module-by-county-by-quarter effects (to allow each module-by-county combination to have arbitrary time trends). The coefficients on price continue to be very close to 1 and are very precisely estimated, and the short-run quantity estimate is roughly -1 (with standard errors ranging from 0.03 to 0.06). The magnitude of the quantity estimate is somewhat smaller in magnitude compared to the corresponding cross-sectional estimate.

To validate our assumption that these estimates do not reflet meaning changes in variety, we report results using log variety as the dependent variable in columns (5) and (6). The point estimates are much smaller in magnitude (compared to the corresponding cross-sectional estimate), though the estimates are sufficiently precise that we can reject the null hypothesis that the point estimates are equal to zero. Interestingly, while the long-run quantity estimates are about 50 percent larger than their short-run counterpart, the long-run variety elasticity is roughly 5 to 6 times larger than the short-run estimate. Thus, our short-run estimates appear to be a reasonable approximation to the "fixed-variety" demand curve, though since the effect on variety is not exactly equal to zero, we will explore sensitivity to adjustments (including estimating variety effect in 2SLS model that does not require variety to actually be held fixed in the short run).

To test whether time-varying unobserved demand shocks may be biasing our short-run estimates, we report robustness checks in Appendix Table $A4$. We find simular results including module-by-quarter-by-DMA fixed effects to absorb any product-specific local demand shocks, and we also rule out that our estimates are driven by quarter-to-quarter product-specific local demand shocks by using an instrument constructed using only stores in the same chain located in other markets. In all cases, the output elasticity is always fairly similar to our main results. Thus, with our short-run price and quantity results and our "long-run" results from pooled cross-sectional regressions, we now have all of the ingredients to estimate the variety effect using formula in equation (16).

4.3 Plug-in estimate of $\tilde{\Lambda}$

Using estimates in Tables 2 and 3, and the illustrative calibrations in Table 1 as a template, we next implement the variety effect formula using different combinations of our reduced-form estimates from the empirical analysis, and we report these results in Table 4. The columns report different reduced-form estimates; column (1) focuses on the short-run and long-run (cross-sectional) estimates that use store and module fixed effects, while column (2) uses analogous estimates that add module-by-county fixed effects. Columns (3) and (4) add module-by-county-by-quarter and store-by-time fixed effect estimates to the short-run estimates. Across all four columns, the plug-in estimates of the variety effect range from 1.1 to 1.2. In other words, the magnitude of estimated variety effect suggests that an exogenous inccrease in variety of 1 percent will inccrease average willingness-to-pay by roughly 1.1 to 1.2 percent.

The four columns of Table 4 focus on the reduced-form effects of the main instrument (leave-me-out average prices) on prices, quantity, and variety, respectively. We can compare these estimates to the variety effect estimates in Kroft et al. (2019a), which reports analogous reduced-form estimates using a different source of exogenous variation: sales taxes. Sales taxes vary across states, and within a state sales taxes also vary across "modules" depending on a state's sales tax exemptions and whether the module is made up of food or non-food products. Kroft et al. (2019a) then generates plug-in estimates of the variety effect based on tax variation. The formula they use differs from equation (17) in two ways. First, unlike prices, sales taxes are not fully salient, and so tax elasticities do not coincide with demand elasticities. The formula therefore needs to adjust for the degree of inattention to taxes. This inattention parameter is estimated by comparing consumer response to prices and taxes as in Chetty, Looney and Kroft (2009). Reassuringly, our estimates of this parameter is very similar to other estimates in the literature. Second, identification of the tax-based variety effect requires the assumption of a flat short-run supply curve (fixed pre-tax prices).[20] Overall, the

[20]The empirical results areconsistent with this assumption. The estimated short-run elasticity of consumer price with respect to taxes is 1.06, consistent with full pass-through of taxes to consumers.

estimates of the variety effect using sales taxes in Kroft et al. (2019a) are in the range 0.2 to 0.7 which is somewhat smaller than the range of estimates 1.1 to 1.2 reported in Table 4. The reduced-form effect of sales taxes on variety and quantity is somewhat attenuated relative to the results in this paper, consistent with sales taxes being less salient. As described in greater detail in Kroft et al. (2019a), consumers responding less to tax changes than to price changes causes producers to adjust variety by less.

4.4 2SLS estimate of $\tilde{\Lambda}$

In addition to the "plug-in" estimates of $\tilde{\Lambda}$, we now show that we can recover an estimate of the variety effect using an alternative approach based on two-stage least squares (2SLS). To see this, note that the variety effect expression in equation (7) motivates the following demand equation:

$$\log(Q_{mrf}) = (-\tilde{\Lambda} * \beta) * \log(J_{mrf}) + \beta * \log(p_{mrf}) + \delta_r + \delta_m + \varepsilon_{mrf} \tag{20}$$

This equation is alsosimilar to the demand equation in the CES-Benassy model in subsection 2.2 since demand depends jointly on variety and prices. The causal effect of prices on demand, holding variety constant, is given by the parameter β. The causal effect of variety on demand corresponds to the parameter $-\tilde{\Lambda} * \beta$. This equation gives another way of interpreting the identification approach described in section 3 – suppose there is an instrument z which affects both prices and variety; then the reduced-form effect $d\log(Q)/dz$ comes from two terms: $(\tilde{\Lambda} * \beta) * d\log(J)/dz$ and $\beta * d\log(p)/dz$. Thus, knowledge of β allows one to recover $\tilde{\Lambda}$ from the reduced-form estimates of the instrument on prices, quantity, and variety as follows:

$$\tilde{\Lambda} \quad = \quad [d\log(Q)/dz - \beta * d\log(p)/dz] * \frac{(1/\beta)}{d\log(J)/dz}$$

This discussion reveals that an alternative approach to identification of the variety effect is to find two instruments that each shift prices and varieties, but by different amounts. For example, suppose there is an instrument that "primarily" affects prices and another instrument that "primarily" affects varieties. Then, the coefficient on prices and varieties can be directly (and jointly) estimated via 2SLS.

The "leave-me-out" instrument based on chain *pricing* was shown above (in 4.1 and 4.2) to affect both prices and varieties. The other instrument we will use in this section is a "leave-me-out" chain *variety* (for each module sold across the stores within a chain). As shown in Appendix Table $A8$, this instrument also strongly predicts the store-level variety (and much more strongly than the leave-me-out price instrument). However, this variety instrument does not predict price as strongly. As a result, these two instruments (leave-me-out price and leave-me-out variety) can be used to jointly identify the coefficients on price and variety in equation (20). This allows us to interpret the coefficient on variety as the genuine variety effect (i.e., causal effect of variety holding average prices constant). As a by-product, it also provides a different way of estimating the fixed-variety (short-run) demand elasticity without using short-run variation.

The 2SLS estimates of equation (20) using the two instruments (leave-me-out average log price and variety) are reported in Table 5, with the different columns corresponding to the same combinations of controls used in Tables 2. For our most restrictive specification (column (2)), the first stage F-statistics range between 9,796 and 17,593 for variety, and between 5,337 and 11,697 for prices, suggesting that the two instruments are jointly very strongly predictive of the two endogenous variables. Across the columns, the implied variety effect estimate (ratio of -1 times the coefficient on log variety divided by the coefficient on average log price) can be compared directly to the "plug-in" estimates of Table 4, and the magnitudes are similar (0.7 to 0.8 instead of 1.1 to 1.2).

Thus, we find that the variety effect estimates are fairly similar using leave-me-out price insturment, sales tax variation, or 2SLS estimates that use both plausibly exogenous variation in both prices and varieties. Part of the similarity in results across these approaches comes from the fact that the *implied* short-run demand elasticity in Table 3 – i.e., the effect of price on quantity holding variety constant – is very similar to the analogous fixed-variety demand elasticity identified using the leave-self-out price and variety instruments, even though the regressions in this section are cross-sectional and do not use any “short-run” variation.

The fact that the “plug-in” estimates are somewhat larger than the 2SLS estimates is perhaps not surprising. A key identifying assumption for the “plug-in” estimate of the variety effect is that the only reason the short-run and long-run (cross-sectional) estimates are different is that the cross-sectional estimates incorporate the effect of endgeonous variety on demand. However, there could be other adjustments over time (besides adjustment to variety). For example, adjustment costs that are larger in the short-run than in the long-run (the LeChatelier principle) will create an upward bias in the estimated variety effect, but the 2SLS estimates would not have this bias (since both the price effect and variety effect are estimated in the same specification). Thus, the somewhat smaller 2SLS variety effect estimates are consistent with meaningful adjustment costs for consumers leading to a more elastic longer-run demand curve, and this would lead to greater change in quantity in the longer term even without any endogenous change in variety.

4.5 Discussion

The results in the previous sections show that the estimates of the variety effect are similar across a range of different approaches: (1) leave-me-out price instrument (as in DellaVigna and Gentzkow 2019), (2) sales taxes (both tax rates and tax exemptions), and (3) using 2SLS with both the chain-level price instrument and chain-level variety instrument. This similarity of estimates across approaches (to different instruments and specifications) is reassuring and suggests that the main variety effect estimates are robust.

We conclude this section with a discussion of the economic significance of the magnitude of the variety effect. Consider the following example: assume that an exogenous 1 percent decrease in variety results in all consumers who previously purchased the (removed) varieties deciding to purchase the outside good. Assuming identical market shares initially, then in this example a 1 percent decrease in variety will reduce quantity demanded by 1 percent (holding price constant). This is identical to an inward shift in the inverse aggregate demand curve by the reciprocal of the price elasticity of demand, or approximately 2 given the short-run estimates in Table 3. Alternatively, we can assume that the individuals who purchased the (removed) varieties instead purchase their next best alternative, which they value 10 percent less on average. In this case, we would calculate an inward shift in the inverse aggregate demand curve by 0.1 percent. Thus, the large variety effect is consistent with many consumers not choosing an alternative variety when their preferred variety is eliminated.

Another way of interpreting magnitude of these results is to compare them to variety effect estimates that would be based on a class of models that link the variety effect to the short-run price elasticity of demand (as in the standard CES model shown above). In the Online Appendix, we report results that impose this equivalence and we tend to find a variety effect that falls within the range of estimates reported earlier. We can also compare our variety effect estimates to other recent estimates from structural trade models that use similar scanner data. Our estimates are similar to the estimates in Handbury and Weinstein (2015), for example.

This paper estimates variety effect averaged across a wide range of product modules, including both food and non-food products. This almost certainly ignores substantial heterogeneity in the variety effect across products. Our identification approach can in principle be applied at finer levels (e.g., combining product modules within narrower category of products). In this spirit, Appendix Table $A6$ reports results that estimate the same reduced-form specifications but only for specific "category" of products: specifically, detergent and soft drinks. The reduced-form estimates for each of these categories can be used to create a category-specific estimates of the variety effect, and we find suggestive evidence of somewhat larger willingness-to-pay for variety in the case of soft drinks as compared to detergent.[21]

Overall, we conclude that the variety effect that we estimate is economically significant and similar across a range of different approaches. Moreover, we have shown in this section that our framework is flexible: it can rely on multiple instruments, or a single instrument (with effects of instrument observed in both "short-run" and "long-run"). In each case, we

[21] This result aligns strongly with the intuition of many seminar participants who had trouble believing a large willingness-to-pay for variety for products such as detergent, but were much more open to idea that consumers have love of variety when it comes to soft drinks.

come to a similar conclusion: consumers prefer variety, so that policies that affect consumer prices (such as taxes and tariffs) that eventually affect variety will have additional effects on consumer surplus above and beyond what is typically estimated when measuring the short-run effects on expenditures.

5 Conclusion

Understanding how changes in product variety affect consumer welfare is critical for a host of questions in economics. In this paper, we develop an empirical approach for valuing changes in product variety and we consider an empirical application. Central to our application are reduced-form estimates of an instrument on prices and quantities where variety is held constant and where variety responds to a change in the instrument, along with estimates of instrument on varieties.

We implement our approach using rich retail scanner data from grocery stores in the US. Our empirical results suggest a reasonably large variety effect. These results are of course specific to this setting, but we see this "proof of concept" as demonstrating applicability to other settings in International Trade and Public Economics. In particular, our approach can be applied in settings where the welfare effect of increasing product variety is of interest. In Kroft et al. (2019a), we consider the welfare effect of taxation, which is of central interest in Public Economics. Standard formulas for measuring the welfare effects of commodity taxes typically assume markets are competitive (Harberger 1964) or markets are characterized by imperfect competition but the number of firms and products are fixed (Auerbach and Hines 2001; Weyl and Fabinger 2013).[22] In differentiated product markets, taxes can distort product variety by reducing firm profitability, and thereby leading to exit for those firms at the margin.

[22]Other papers in this literature include Seade (1987), Stern (1987), Myles (1989), Besley (1989), Delipalla and Keen (1992), Anderson, de Palma and Kreider (2001a, 2001b), Auerbach and Hines (2001), Weyl and Fabinger (2013) and Gillitzer, Kleven and Slemrod (2015). These papers typically assume a specific form of firm competition and impose specific structure on consumer preferences, and some of these papers focus purely on the short-run equilibrium holding the number of varieties fixed. Below we allow for both differentiated products and free entry in deriving our welfare expressions, without having to specify the form of imperfect competition (e.g. Bertrand vs Cournot).

In this context, pass-through alone (as in Weyl and Fabinger 2013) is no longer a sufficient statistic for the welfare effects of taxation, instead we also need a measure of the variety effect. We derive a new formula for the marginal welfare gain from increasing commodity taxes in a general model of imperfect competition, allowing for salience effects following Chetty, Looney and Kroft (2009), covering a wide range of market conduct when variety is endogenous. While the standard formula emphasizes the response of total output to the tax due to a fiscal externality, our formula shows additionally that one needs to account for the variety effect which arises since taxes distort product variety. Interestingly, the impact on welfare coming through a change in variety is ambiguous and depends on whether variety is excessive or insufficient at the prevailing equilibrium (Mankiw and Whinston 1986). Finally, we show how our new formula connects to existing formulas for the welfare effects of taxes under imperfect competition.

Finally, one can apply our analysis of the variety effect developed here to study the impact of tariffs on welfare and other government policies that result in changes in equilibrium varieties, such as price controls (e.g., rent control) and mergers, taking into account the variety effect in addition to the price effect.

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Table 1: Illustrative Calculations of the Variety Effect Based on Reduced-Form Estimates

	(1)	(2)	(3)	(4)
Cross-sectional reduced-form estimates (endogenous J)				
Price response, dlog(p)/dt	1.0	1.0	1.0	1.0
Quantity response, dlog(Q)/dt	-1.0	-1.6	-1.6	-1.3
Variety response, dlog(J)/dt	-0.6	-0.6	-0.3	-0.6
Short-run reduced-form estimates (fixed J)				
Price response, dlog(p)/dt |J	1.0	1.0	1.0	1.0
Quantity response, dlog(Q)/dt |J	-1.0	-1.0	-1.0	-1.0
Variety effect "plug-in" estimate				
Variety Effect Parameter	0.0	1.0	2.0	0.5

Notes: This table reports calibration of the variety effect formula in main text. All columns report illustrative calibrations for different scenarios making different assumptions about short-run and long-run reduced-form estimates. The first column illustrates the scenarios where consumers do not value variety. The remaining columns show how the magnitude of the variety effect varies with the reduced-form estimates.

Table 2 : Cross-Sectional Reduced-Form Estimates

Dependent Variable:	Prices		Quantity		Variety	
	(1)	(2)	(3)	(4)	(5)	(6)
Leave-self-out average log(*p*)	0.997	0.915	-1.881	-1.531	-0.687	-0.478
	(0.000)	(0.006)	(0.115)	(0.075)	(0.076)	(0.053)
Specification:						
Store, Time, Module fixed effects	y	y	y	y	y	y
Module x County fixed effects		y		y		y
N (observations)	14,057,165	14,057,165	14,057,165	14,057,165	14,057,165	14,057,165
N (modules)	198	198	198	198	198	198
N (stores)	8,653	8,653	8,653	8,653	8,653	8,653
N (chains)	70	70	70	70	70	70
R^2	0.894	0.946	0.855	0.944	0.871	0.943

Notes: All dependent variables are measured yearly. The Nielsen Retail Scanner data is restricted to modules above the 80th percentile of the national distribution of sales. The independent variable is the leave-self-out retail chain average of log price (averaging across all other stores in the chain). All coefficients are equally-weighted linear combinations of nine coefficients, one cross-ectional estimate for each year from 2006 to 2014. Standard errors are clustered at the chain-module level and estimated using Seemingly Unrelated Regression.

Table 3: Short-Run Reduced-Form Estimates

Dependent Variable:	Prices		Quantity		Variety	
	(1)	(2)	(3)	(4)	(5)	(6)
Leave-self-out average log(p)	0.994	0.964	-1.029	-1.022	-0.109	-0.145
	(0.000)	(0.002)	(0.063)	(0.029)	(0.038)	(0.023)
Specification:						
Store, Time, Module fixed effects	y	y	y	y	y	y
Module × Store fixed effects	y	y	y	y	y	y
Module × Time fixed effects	y	y	y	y	y	y
Module × Time × County fixed effects		y		y		y
Store × Time fixed effects		y		y		y
N (observations)	55,536,376	55,536,376	55,536,376	55,536,376	55,536,376	55,536,376
N (modules)	198	198	198	198	198	198
N (stores)	8,653	8,653	8,653	8,653	8,653	8,653
N (chains)	70	70	70	70	70	70
N (quarters)	36	36	36	36	36	36
R^2	0.915	0.951	0.961	0.988	0.966	0.983

Notes: All dependent variables are measured at year-quarter, and the unit of observation is store-by-module. The period covered is 2006-2014. See notes to Table 2 for more details.

Table 4: Plug-In Estimates of the Variety Effect

	Leave-self-out reduced-form estimates			
Cross-sectional reduced-form estimates (Table 2)	Odd # cols.	Even # cols.	Odd # cols.	Even # cols.
Short-run reduced-form estimates (Table 3)	Odd # cols.	Odd # cols.	Even # cols.	Even # cols.
	(1)	(2)	(3)	(4)
Cross-sectional estimates (endogenous J)				
Price response, dlog(p)/dz	1.00	0.92	1.00	0.92
Output response, dlog(Q)/dz	-1.88	-1.53	-1.88	-1.53
Variety response, dlog(J)/dz	-0.69	-0.48	-0.69	-0.48
Short-run estimates (fixed J)				
Price response, dlog(p)/dz |J	0.99	0.99	0.96	0.96
Quantity response, dlog(Q)/dz |J	-1.03	-1.03	-1.02	-1.02
Variety effect "plug-in" estimate				
Variety Effect Parameter	1.193	1.178	1.131	1.105

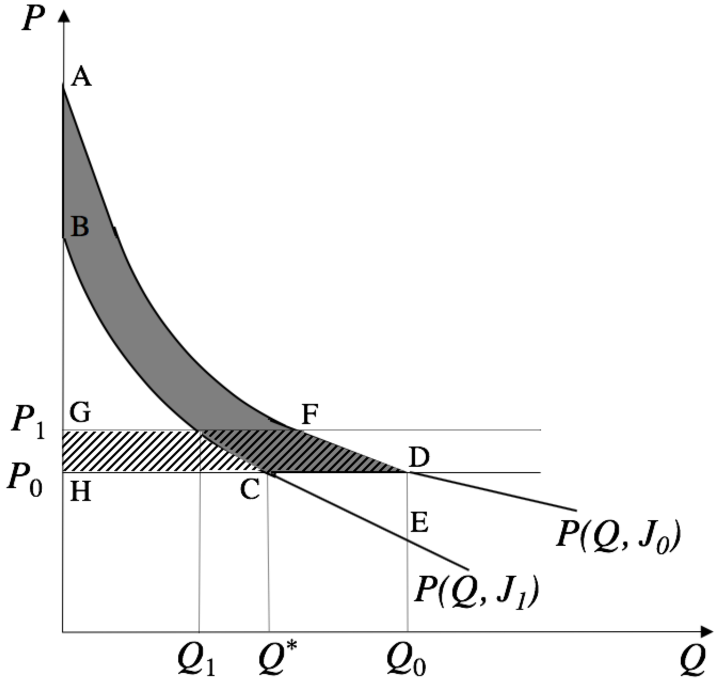
Notes: This table reports "plug-in" estimates of the variety effect using different combinations of estimates from Tables 2 and 3.

Table 5: 2SLS Estimates of Variety Effect and Short-Run Demand Elasticity

Dependent Variable:	log(Quantity)	
	(1)	(2)
log(Variety)	1.007	0.840
	(0.019)	(0.018)
log(Price)	-1.193	-1.234
	(0.059)	(0.056)
Implied variety effect estimate	0.845	0.681
Specification:		
Store, Time, Module fixed effects	y	y
Module x County fixed effects		y
N (observations)	14,057,165	14,057,165
N (modules)	198	198
N (stores)	8,653	8,653
N (chains)	70	70
R^2	0.914	0.962

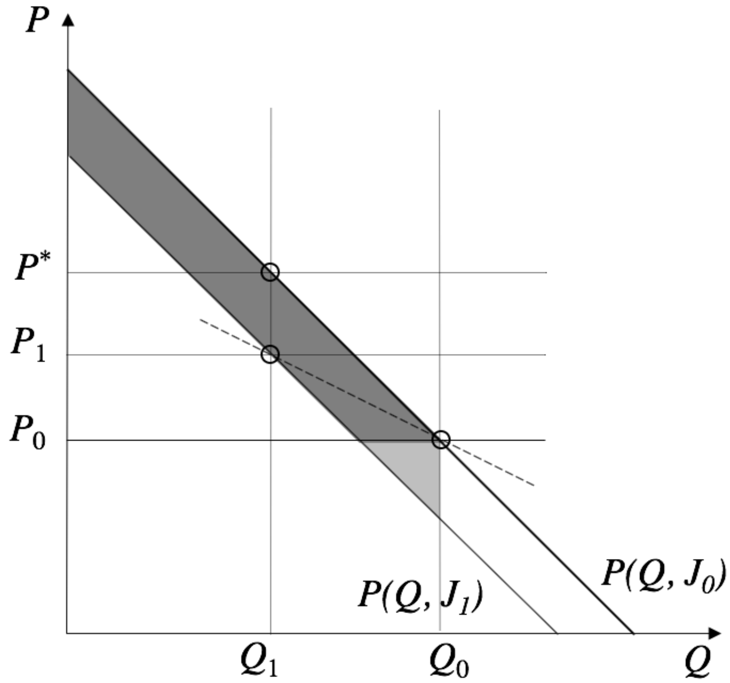
Notes: All dependent variables are measured yearly and enter in logs. The Retail Scanner data is restricted to modules above the 80th percentile of the national distribution of sales. The two endogenous variables (Price and Variety) are instrumented with Leave-self-out average log price and variety. All coefficients are equally-weighted linear combinations of nine coefficients -- one for each year from 2006 to 2014. Standard errors are clustered at the chain-module level and estimated using a stacked 2SLS model (that stacks each cross-sectional 2SLS regression for each year), analogous to Seemingly Unrelated Regression.

Figure 1: Price Effect and Variety Effect



Notes: This figure shows graphically the the result of a decrease in variety (from J_0 to J_1). The price effect is the hatched pattern area $DFGH$, which is formed by the base of the pre-existing quantity (Q_0) and the change in prices (from P_0 to P_1). The dark shaded area is the variety effect which is the area between the two aggregate demand curves.

Figure 2: Variety Effect Under Parallel Demands



Notes: This figure shows graphically the variety effect in the case of parallel aggregate demand curves. In this case, the variety effect is the area of the shaded parallelogram. The demand curves are drawn to be linear for ease of interpretation. The variety effect is equal to the dark shaded area above P_0. In our empirical implementation, we measure the variety effect as this area plus the light shaded triangle below P_0 which is approximately zero for a small change in variety according to Corollary 1.